The Small Open Economy New Keynesian Phillips Curve: Specification, Structural Breaks and Robustness*

Juan Carlos Aquino†

Central Reserve Bank of Peru

This version: July 15, 2019

Abstract

This paper empirically assesses the concern on whether the slope of the Phillips with respect to the output gap has decreased (i.e. the Phillips curve has "flattened"). We derive a generalized lag-augmented version of the New-Keynesian Phillips Curve for a small open economy (Galí and Monacelli, 2005) in order to specify a semi-structural estimation equation. For the Peruvian economy, such equation is estimated *via* the Generalized Method of Moments for the Inflation-Targeting regime (January 2002 - March 2019) and the post-crisis (January 2008 - March 2019) periods. We found that the slope parameter has remained stable for both estimation periods. Moreover, the expectation channel has gained more relevance for the post-crisis period, a result that is consistent with a lower persistence of inflation dynamics. Our results are also consistent with the presence of full price indexation across estimation samples.

JEL Classification: C22, C51, E31.

Keywords: New-Keynesian Phillips curve, small open economy, Generalized Method of Moments.

*I thank Marco Vega, Paul Castillo, Youel Rojas, and Raimundo Chirinos for comments and suggestions, and María Gracia Ramos who additionally provided support with data. The usual disclaimer applies.

†Research Division. Mail address: Jirón Santa Rosa 441-445, Lima, Peru. Postal code 15001. E-mail address: aquinojc0806@gmail.com.

1 Introduction

One of the most important ways in which monetary policy affects inflation is through its effects on economic activity. Such channel is usually represented by a (positive) relation between inflation and a measure of "inflationary pressures" known as the Phillips curve and inspired by the work of Phillips (1958). For the case of Peru, Figure 1 depicts the quarterly evolution of the four-quarter core inflation and the cyclical component of GDP (a.k.a. output gap) since 1999. It can be noticed that, since the adoption of the Inflation Targeting regime in 2002, such relation has apparently prevailed until 2013 (gray shaded area). However, it can also be noticed that the same relation has apparently eroded from 2014 onwards (dark gray shaded area) which naturally raises concerns about the effectiveness of monetary policy.

From a technical point of view, the previous discussion is often organized in terms of the New-Keynesian Phillips Curve (henceforth, NKPC)

$$\pi_t = \beta E_t \pi_{t+1} + \kappa(y_t - g_t) + \varepsilon_t$$

(see Clarida et al., 1999) where π_t denotes the inflation rate, $E_t\pi_{t+1}$ denotes the expected inflation rate, y_t denotes the output level, g_t denotes the potential output level, β and κ are positive constants, and ε_t is a random disturbance term. In this regard, the recent episodes of economic contraction and lack of deflationary pressures led to a concern on whether the Phillips curve has "flattened" or, equivalently, the slope parameter κ has decreased.

In this paper, we perform a semi-structural estimation of a NKPC for Peru in order to answer whether or not the Phillips curve has flattened. Our approach incorporates some novel features. First, our reduced-form specification (and model-based sign restrictions) arises from our derivation of a hybrid (lag-augmented) version of the small open economy NKPC by Galí and Monacelli (2005) in order to account for inflation persistence. Second, our extension is compatible with monthly data available for the entire Inflation Targeting regime adopted by the Central Reserve of Peru in 2002. Third, our estimates are obtained *via* the Generalized Method of Moments (GMM) developed by Hansen (1982) and the moment-selection criteria proposed by Andrews (1999). Finally, we report our results for two estimation samples, 2002-2019 (entire Inflation Targeting regime) and 2008-2019 (post-crisis period), in order to check for parameter stability.

Our findings are summarized as follows. First, our estimates support the theory-based sign restrictions. Second, for both estimation samples, the slope parameter has remained stable and thus the corresponding channel of monetary policy is unaltered. Third, compared to the full Inflation-Targeting regime, for the post-crisis period the expectation channel has gained more relevance and this finding is consistent with a lower persistence of inflation. Finally, our results are consistent with the presence of full price indexation for both estimation samples.

The rest of this paper is organized as follows. Section 2 provides a (non-exhaustive) review of related literature, with a special emphasis on alternative derivations of the NKPC. Section 3 presents the theoretical framework that leads to the (semi-structural) specification to be employed in the estimation process. Section 4 briefly describes both the GMM estimator and our testable hypotheses of interest. Section 5 reports our estimation results. Section 6 concludes.

2 Related Literature

The starting point is given by the contributions of Calvo (1983) and Yun (1996), which led to the (benchmark) small open economy version of the New-Keynesian Phillips Curve by Galí and Monacelli (2005)

$$\pi_{H,t} = \beta E_t \pi_{H,t+1} + \kappa_\upsilon \tilde{y}_t \tag{1}$$

where $\beta \in (0,1)$ is the discount factor and $\kappa_\upsilon > 0$ is expressed in terms of the deep parameters of the model.

Etro and Rossi (2015) study a New-Keynesian model for the case of a small and fixed number of firms competing in prices *à la* Bertrand and with the corresponding slope being lower than those for the case of atomistic firms.

de Gregorio et al. (2007) extend the benchmark model study oil and food price shocks. Portillo and Zanna (2015) develop a tractable small open-economy model to study the first-round effects of international food price shocks in developing countries. Okano (2014) derives a NKPC for the Producer Price Index inflation within a setting that distinguishes between tradable and non-tradable goods and extends Galí and Monacelli (2005) by adding up a term (besides the output gap) that captures the non-tradable price disparity between countries. Kamber et al. (2016) develops an extension with tradeable and non-tradeable goods in an approach similar to the one developed by Adolfson et al. (2007). Popescu (2014) derives a forward-looking NKPC for the headline inflation rate that also includes the current and expected exchange rate.

In Fasolo (2014), firms in the tradable and non-tradable sectors of the domestic economy rent capital and labor from households to produce goods. They set prices in a Calvo style, with an exogenous probability of optimizing prices. Bouakez and Eyquem (2015) introduced the differential effect of the real exchange rate.

Craighead (2014) incorporates a search-and-matching model of the labor market into a small open economy model with nominal rigidities. Such feature allows the behavior of tradable and non-tradable sector unemployment rates to be studied under alternative monetary rules. Unlike the customary off-the-shelf Galí-Monacelli-type model to several heterogeneous countries, a major departure from this benchmark model is developed by Cacciatore et al. (2016) who introduce richer micro-level producer dynamics and search-and-matching frictions in labor markets for the case of South Korea.

Another extension is provided by Monacelli (2005) who incorporates price-setting retailers and introduces incomplete exchange rate pass-through on import prices. Such extension is proposed in order to obtain an aggregate supply curve for import prices and therefore to account for empirical evidence on the dynamics of import pricing for some OECD economies reported by Campa and Goldberg (2005).

On the other hand, the reasons to include a framework with incomplete asset markets are stresses out by Alonso-Carrera and Kam (2016) as market incompleteness exacerbates the domestic-inflation and output-gap monetary-policy trade-off.

In a major extension, Justiniano and Preston (2010) extend the framework by Galí and Monacelli (2005) by allowing habit formation in consumption, incomplete asset markets, and price indexation to past inflation. It delivers a NKPC of the form

$$\pi_{H,t} - \delta\pi_{H,t-1} = \theta_H^{-1}(1-\theta_H)(1-\theta_H\beta)mc_t + \beta E_t\left\{\pi_{H,t+1} - \delta\pi_{H,t}\right\}$$

where the marginal costs are given by $mc_t = \varphi y_t - (1+\varphi)\varepsilon_{a,t} + \alpha s_t + \sigma(1-h)^{-1}(c_t - hc_{t-1})$, $\varepsilon_{a,t}$ is a productivity shock and c_t is private consumption. Again, the reader is referred to Justiniano and Preston (2010) for an interpretation of the structural parameters.

On the empirical side, Anguyo et al. (2017) embed the model by Justiniano and Preston (2010) into an estimated regime-switching dynamic stochastic general equilibrium model for monetary policy analysis and forecasting purposes for the case of Uganda, in a model that exhibits financial frictions.

Sánchez (2018) employs a Markov-switching Dynamic Stochastic General Equilibrium (MS-

DSGE) model to identify regime switches in the driven mechanisms of the Colombian economy for the 1990-2014 period. Kolasa and Rubaszek (2018) compare the forecasting performance of the small open-economy DSGE model by Justiniano and Preston (2010), among others, against a closed-economy benchmark using data for Australia, Canada and the United Kingdom. Finally, Lie (2018) estimates a Dynamic Stochastic General Equilibrium (DSGE) model for Indonesia in order to study the effect of (official) inflation-target adjustments on fluctuations at the aggregate level. The framework is primarily based on Justiniano and Preston (2010) albeit with two extensions: i) the inclusion of potential adjustments in the central bank's inflation target and ii) a money-holding friction due to a cash-in-advance (CIA) constraint is assumed.

Kamber et al. (2015) presents a structural model for policy analysis and forecasting that includes several Phillips Curves for domestic goods, tradable goods and exports.

Another branch of the literature emphasized the use of semi-structural models. Dungey et al. (2014) semi-structural approach which emphasizes the variation of the real excange rate. Kichian and Rumler (2014) employ a semi-structural approach for forecasting purposes based on Gali and Monacheli. For example, Yeh (2017) employs a semi-structural NKPC augmented by asset prices to study the adoption of alternative policy targets.

Recently, some of these theoretical extensions have been combined with econometric methodologies. For instance, Kavtaradze (2014) employs a hybrid NKPC nested within a time-varying parameter (TVP) framework that incorporates both forward-looking and backward-looking components. Other studies have relied on Markov-switching models, like Debortoli and Nunes (2014) who embeds a NKPC to reflect changes in the policy maker's preferences. Other examples are Kriwoluzky et al. (2015) who also use a Markov-switching model that allows for changes in parameters for the case of Greece. Davis et al. (2017) combine the Markov-switching with a standard NKPC that also includes the real exchange rate. For the case of South Africa, Balcilar et al. (2017) estimate a Markov-switching Dynamic Stochastic General Equilibrium model for the 1989-2014 period that exhibits a lag-augmented version of the NKPC with complete financial markets.

3 Theoretical Framework

The starting point towards our econometric model specification is the New-Keynesian framework for a small open economy by Galí and Monacelli (2005) and the reader is referred for further details to Galí (2015, Chapter 8) which is the exposition we borrowed the notation from. Specifically, $\beta \in (0,1)$ is the domestic households' discount factor, $\upsilon \in [0,1]$ represents the share of foreign goods in domestic composite consumption and therefore can be interpreted as a measure of openness, $\eta > 0$ measures the substitutability between domestic and foreign goods, $\epsilon > 1$ denotes the elasticity of substitution between varieties produced domestically, $\sigma > 0$ is the inverse of the intertemporal elasticity of substitution, $\varphi > 0$ is the inverse of the (real) wage elasticity of domestic households' labor supply, $1-\alpha \in (0,1)$ represents the elasticity of domestic output with respect to labor and $\theta \in (0,1)$ measures the fraction of domestic firms that cannot set new prices each period.

Our extension to the previous framework is described as follows: producers who are not allowed to reset their prices rather index them to the last q realizations of the domestic inflation rate $\pi_{H,t-1}$, $\pi_{H,t-2}$, ... and $\pi_{H,t-q}$ with non-negative coefficients ρ_1, ρ_2, ... and ρ_q, respectively. Following Sbordone (2005) and Magnusson and Mavroeidis (2014), it is easy to show that our extension leads to the following hybrid New-Keynesian Phillips Curve for the domestic inflation rate $\pi_{H,t}$

$$\pi_{H,t} = \frac{\rho(L) - \beta \rho_{\Delta}(L)}{1+\beta\rho_1}\pi_{t-1} + \frac{\beta}{1+\beta\rho_1}E_t\pi_{H,t+1} + \kappa'_{\upsilon}\tilde{y}_t \tag{2}$$

where the polynomials $\rho(L) = \rho_1 + \rho_2 L + \ldots + \rho_q L^{q-1}$ and $\rho_{\Delta}(L) = \rho_2 + \rho_3 L + \ldots + \rho_q L^{q-2}$ are expressed in terms of the lag operator L. Also, the slope of (2) with respect to the output gap $\tilde{y}_t$ is given by $\kappa'_{\upsilon} = \lambda'(\sigma_{\upsilon} + \frac{\varphi+\alpha}{1-\alpha}) > 0$ where the terms $\lambda' = \frac{(1-\theta)(1-\beta\theta)}{\theta(1+\beta\rho_1)}\Theta$, $\Theta = \frac{1-\alpha}{1-\alpha+\alpha\epsilon}$, $\sigma_{\upsilon} = \sigma\Phi$, $\Phi = \frac{1}{1+\upsilon(\varpi-1)}$ and $\varpi = \sigma\eta + (1-\upsilon)(\sigma\eta - 1)$ are all positive in the parameter space.

For $\beta \approx 1$ and $q = 3$ we obtain

$$\pi_{H,t} = \frac{\rho_1 - \rho_2}{1+\rho_1}\pi_{H,t-1} + \frac{\rho_2 - \rho_3}{1+\rho_1}\pi_{H,t-2} + \frac{\rho_3}{1+\rho_1}\pi_{H,t-3} + \frac{1}{1+\rho_1}E_t\pi_{H,t+1} + \kappa'_{\upsilon}\tilde{y}_t \tag{3}$$

which is expressed in terms of the deep parameters in $(\rho_1, \rho_2, \rho_3, \upsilon, \eta, \epsilon, \sigma, \varphi, \alpha, \theta)$. Some comments are in order. First, (3) imposes no sign restriction on the coefficients associated to either $\pi_{H,t-1}$ or $\pi_{H,t-2}$. Second, the coefficient associated to $\pi_{H,t-3}$ is allowed to be greater than or equal to zero. Third, the coefficients corresponding to the expected domestic inflation $E_t\pi_{H,t+1}$ and the output gap $\tilde{y}_t$ are both strictly positive and provide testable hypotheses. Fourth, the coefficients associated to the lagged and expected (domestic) inflation add up to 1 (i.e. there is full price indexation) and this feature also provides a testable hypothesis. Finally, for the case of no indexation ($\rho_1 = \rho_2 = \ldots = \rho_q = 0$), equation (2) leads to the canonical representation of the New-Keynesian Phillips Curve in Galí (2015, Chapter 8, equation 37).

4 Empirical Strategy

4.1 Generalized Method of Moments (GMM) Estimator

The equilibrium New-Keynesian Phillips Curve (3) underlies the following reduced-form equation for estimation purposes:

$$\pi_{H,t} = c_0 + c_1\pi_{H,t-1} + c_2\pi_{H,t-2} + c_3\pi_{H,t-3} + c_{exp}\pi_{H,t+1} + c_{gap}\tilde{y}_t + u_t, \tag{4}$$

where c_0 is a constant term, c_i is the coefficient of the i-th lag of the domestic inflation $\pi_{H,t-i}$ ($i = 1, 2, 3$) and is intended to capture inflation inertia, c_{exp} is the coefficient of the future domestic inflation $\pi_{H,t+1}$ and is intended to capture the expectation channel,[1] c_{gap} is the coefficient of the output gap $\tilde{y}_t$ (i.e. the "slope" of the New-Keynesian Phillips Curve) and u_t contains the forecasting error $\pi_{H,t+1} - E_t\pi_{H,t+1}$.

Let $x_t \equiv (\pi_{H,t}, \pi_{H,t-1}, \pi_{H,t-2}, \pi_{H,t-3}, \pi_{H,t+1}, \tilde{y}_t)$ contain the variables involved in (4) and let $c \equiv (c_0, c_1, c_2, c_3, c_{exp}, c_{gap})$ contain the reduced-form coefficients in (4). Also, let

$$m(x_t; c) \equiv \pi_{H,t} - (c_0 + c_1\pi_{H,t-1} + c_2\pi_{H,t-2} + c_3\pi_{H,t-3} + c_{exp}\pi_{H,t+1} + c_{gap}\tilde{y}_t) \tag{5}$$

denote the forecasting error u_t and c^0 denote the coefficient vector of the data generating process. Under rational expectations, the equation (4) evaluated at $c = c^0$ implies that the unconditional expectation of the forecasting error u_t equals zero (i.e. $E\left[m(x_t; c^0)\right] = 0$). Also, under rational expectations such forecasting error is uncorrelated to any variable in the agents' information set. Let $z_{j,t}$ ($j = 1, \ldots, p$) represent such variable. Then, the previous description leads to p moment conditions for p (instrumental) variables $\{z_{1,t}, \ldots, z_{p,t}\}$ in the information set the form $E\left[z_{j,t}m(x_t; c^0)\right] = 0$ ($j = 1, \ldots, p$) or, compactly,

$$E\left[Z_t m(x_t; c^0)\right] = 0 \tag{6}$$

where $Z_t = [z_{1,t} \ \ldots \ z_{p,t}]'$ is the vector of instrumental variables. Since the vector c contains six coefficients, we restrict to the case of over-identification by assuming $p > 6$. The Generalized Method of Moment (GMM) estimator by Hansen (1982) estimates c^0 by finding the c that

[1] Although there exists an available series on agents' expectations since the beginning of the Inflation Targeting regime, such information is not employed for two reasons. First, it provides agents' expected headline inflation, whereas our model is posed in terms of domestic inflation. And second, it consists of a 12-month ahead expectation, whereas our model is posed in terms of a one-month ahead expectation.

makes the sample analogue of (6) as close to zero as possible through the use of a weighing matrix. Specifically, for a sample of size T the GMM estimator $\hat{c}_{GMM}$ minimizes

$$\mathcal{L}_{GMM}(c) \equiv g_T'(c)\,\hat{N}_u^{-1} g_T(c)\,, \text{ with } g_T(c) = T^{-1}\sum_{t=1}^{T} g_t(c), \tag{7}$$

where $g_t(c) = Z_t m(x_t; c)$ and $\hat{N}_u \xrightarrow{p} N_u = \texttt{lim}_T \left[\texttt{Var}[\sqrt{T} g_T(c^0)]\right] \equiv \texttt{AVar}[\sqrt{T} g_T(c^0)]$.

The reader should notice that the estimator is based on the assumption that the vector of instruments Z_t satisfies the over-identifying conditions (6). For large T and under the null hypothesis that such over-identifying restrictions are all valid, the Sargan's J-statistic $J_T(\hat{c}_{GMM}) \equiv T g_T'(\hat{c}_{GMM})\,\hat{N}_u^{-1} g_T(\hat{c}_{GMM})$ is chi-squared distributed with $p-6$ degrees of freedom and cumulative distribution denoted by F. Let $\tilde{\alpha}$ denote the chosen significance level. Therefore, we reject the null hypothesis of over-identification if the calculated p-value $1 - F(J_T(\hat{c}_{GMM}))$ is greater than $\tilde{\alpha}$ and cannot reject it otherwise.

Our consistent moment selection follows Andrews (1999) as it involved a search along vectors Z_t that contain a constant and instruments within the set $\{\pi_{H,t-k}, \tilde{y}_{t-k}\}_{k=1}^{k_{max}}$ and GMM analogues of the Bayesian, Akaike and Hannan-Quinn Information Criteria for moment selection were employed. We refer to them as GMM-BIC, GMM-AIC and GMM-HQIC and define them by

$$\begin{aligned}
\text{GMM-BIC} &: \quad MSC_{BIC,T}(Z_t) = J_T(Z_t) - (p-6)\log T;\\
\text{GMM-AIC} &: \quad MSC_{AIC,T}(Z_t) = J_T(Z_t) - 2 \times (p-6)\log T;\\
\text{GMM-HQIC} &: \quad MSC_{HQIC,T}(Z_t) = J_T(Z_t) - 2.01 \times (p-6)\log\log T;
\end{aligned}$$

where Z_t is a vector containing p instruments, and log denotes natural logarithm.

4.2 Data

Our theoretical framework implies an econometric specification involving only two variables: the period-to-period domestic inflation rate and the output gap. We consider monthly data for the period from January of 2002 to March of 2019 for two reasons. First, a (New-Keynesian) Phillips Curve is one of the key ingredients of the Inflation Targeting (IT) regime that the Central Reserve Bank of Peru (BCRP) adopted in 2002, in which monetary policy decisions are made on a monthly basis and there is available data for the same period and frequency. Second, the use of monthly data implies 207 observations from January of 2002 to March of 2019 which, unlike the 69 quarterly observations for the same time span, provides more data variability in the estimation process. The latter feature will allow our statistical inference to rely on large-sample distributions as we assume that they fairly approximate finite-sample distributions of tests statistics.

Our theoretical framework is also explicit regarding the variables to include and the transformations to perform. Also, all raw variables were obtained from the Central Reserve Bank of Peru's database. The variable representing domestic inflation $\pi_{H,t}$ is given by $100\times\Delta\log(\text{IPC\_h})$, the (natural) logarithm of the domestic component of the monthly Consumer Price Index in first differences. On the other hand, a proxy for the output gap $\tilde{y}_t$ is given by $100 \times \text{output\_gap}$, the difference between the (natural) logarithm of the seasonally-adjusted monthly Economic Activity Index and its Hodrick-Prescott (HP) filter trend.[2] It is worth to mention that our seasonal adjustment made use of the automatic mode of the programs TRAMO and SEATS which implement the methodology proposed by Gomez and Maravall (1994) and available on the Bank of Spain's website.

4.3 Specification and testable hypotheses

Therefore, a semi-structural specification based on (3) and suitable for estimation is

$$\begin{aligned}\Delta\log(\text{IPC\_h}) = c_0 + c_1 \times \Delta\log(\text{IPC\_h(-1)}) + c_2 \times \Delta\log(\text{IPC\_h(-2)})\\ + c_3 \times \Delta\log(\text{IPC\_h(-3)}) + c_{exp} \times \Delta\log(\text{IPC\_h(+1)}) + c_{gap} \times \text{output\_gap} + \text{u} \quad (8)\end{aligned}$$

where u is an error term that contains preference and technology shocks and domestic inflation forecasting errors as we include the actual future domestic inflation rate $\Delta\log(\text{IPC\_h(+1)})$ instead of its conditional expectation. In addition to the usual tests of significance, there are three hypotheses we are interested in:

1. $H_0 : c_{exp} \leq 0$ against $H_1 : c_{exp} > 0$ (expectations matter in the NKPC),
2. $H_0 : c_{gap} \leq 0$ against $H_1 : c_{gap} > 0$ (positive slope of the NKPC), and

[2] For this purpose, the standard value of the smoothing parameter ($\lambda = 14,400$) was employed. Also, in order to mitigate the end-point bias, the calculations also included the ARIMA forecasts from April of 2019 to December of 2019. Finally, it is worth to emphasize that, unlike $\tilde{y}_t$, the use of filtered data implies that the error term u now also contains the irregular components of the flexible-price output level such as preference and technology shocks.

3. $H_0 : c_1 + c_2 + c_3 + c_{exp} = 1$ (full indexation) against $H_1 : c_1 + c_2 + c_3 + c_{exp} \neq 1$.

It is in this regard that rejecting the null hypothesis in 1 would support the alternative hypothesis that the expectations are relevant for domestic inflation dynamics. A similar description applies to the hypotheses in 2 regarding the slope of the New-Keynesian Phillips Curve and thus the effect of the output gap. Finally, the null hypothesis in 3 is consistent with full price indexation as specified by our theoretical model.

5 Results

5.1 Unit Root Testing

As it is customary, the detection of unit roots becomes relevant for the specification of our empirical model. This so happens because all the variables included in (8) are assumed to be stationary. For this reason, the unit root tests by Dickey and Fuller (1979), Said and Dickey (1984) and Phillips and Perron (1988) are reported in Table 1. We reject the null hypothesis that output_gap contains a unit root and cannot reject the null hypothesis that log(IPC_h) contains a unit root. Both results hold for all conventional significance levels (1%, 5% and 10%) and regardless of the specification of the deterministic component. The same results are obtained for the efficient test developed by Elliott et al. (1996) in Tables 2 and 3, and for the class of M-test by Ng and Perron (2001) in Table 4 that overcome a series of well known limitations involving the power loss of unit root tests against local alternatives.

Nevertheless, it can be noticed in Figure 2 that log(IPC_h) seems to exhibit a trend shift. A similar pattern is observed for output_gap in Figure 6. According to Perron (1989), such abrupt shifts distort conventional unit root tests and lead to an over acceptance of the unit root hypothesis. For this reason, Table 5 reports the unit root tests proposed by Perron and Rodríguez (2003) which allow for the presence of a structural change. That is, a trend shift is allowed and “controlled” in a robust fashion while testing for unit roots. Once again, we reject the null hypothesis that output_gap contains a unit root and cannot reject the null hypothesis that log(IPC_h) contains a unit root at all conventional significance levels.

However, and by construction, the tests by Perron (1989) pre-assume the existence of a break with non-trivial effects on its power. Moreover, a detected break date can turn out to be spurious. For this reason. the tests by Cavaliere et al. (2011) pre-test for the existence of a break in the trend function. At the 5% level of significance, we reject the null that output_gap as has a unit root with a structural break. Also, at the same level of significance we cannot reject the null that log(IPC_h) has a unit root with a structural break.

5.2 GMM Estimation and Hypothesis Testing

Table 7 summarizes our estimates of the coefficients in equation (8) for two estimation periods and several instrument sets. Columns I, II and III contain estimates for the period from January of 2002 to March of 2019 (i.e. from the beginning of the Inflation Targeting regime) whereas columns IV, V and VI contain estimates for the period from January of 2008 onwards (i.e. consistent with the last financial crisis) since the univariate unit root test by Perron and Rodríguez (2003) estimates a structural break for the (log of) domestic prices as occurring during January of 2008. The Generalized Method of Moments (GMM) estimator by Hansen (1982) was employed for all of the equations. Also, for all cases, the effective numbers of observations are lower that those implied by the original time span because of the lagged variables being employed as regressors and/or instruments. For each estimation period, we set $k_{max} = 7$ (i.e. the maximum lag used as an instrument) and an exhaustive search for instruments was performed. Results for instrument sets exhibiting the three lowest moment selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are also reported in Table 8. In our search, we filtered out any instrument vector such that the null hypothesis of over-identification is rejected. It is worth to notice that for both estimation periods, each moment selection criterion monotonically decreases (from III to I and from VI to IV). This reflects that the different bonus terms (that reward selection vectors that utilize more moment conditions) have no impact on the corresponding moment-selection criterion and therefore the problem of moment selection reduces to minimize the Sargan's J-statistic with respect to the instrument vector.

On the one hand, from column I it can be asserted that, regarding the sign-unrestricted coefficients, the lagged domestic inflation Δlog(IPC_h(-1)) is significant at any of the conventional significance levels (either 1%, 5% or 10%) and has a positive marginal effect that equals 0.32. On the contrary, neither Δlog(IPC_h(-2)) nor Δlog(IPC_h(-3)) are individually significant at any of the conventional significance levels. The point estimate of the marginal effect of Δlog(IPC_h(-2)) is negative (a possibility captured by the theoretical model). Also, we cannot reject the null hypothesis that Δlog(IPC_h(-3)) is not significant. On the other hand, regarding the sign-restricted coefficients, we reject that the domestic inflation expectation (output gap) has a lower-than-or-equal-to-zero effect at the 10% significance level and conclude that there exists a positive and significative effect. Such conclusion is reflected by a one-sided p-value lower than 0.10. Finally, we cannot reject the null hypothesis that there is full indexation at the 10% significance level, which is reflected by a two-sided p-value that equals 0.13 (greater than a conservative 0.10). A similar analysis applies to both columns II and III.

For the post-crisis period, from column IV it is found that, regarding the sign-unrestricted coefficients, the lagged domestic inflation Δlog(IPC_h(-1)) is again significant at any conventional significance level although the point estimate of the marginal effect now equals 0.17 which is lower than 0.32 for the full-sample estimation. From a standpoint based on a theoretical framework, this sheds light on what structural feature might be driving the change in the inflation dynamics. Namely, after the financial crisis this is consistent with a lower fraction of firms indexing their prices to the previous domestic inflation. Again, neither Δlog(IPC_h(-2)) nor Δlog(IPC_h(-3)) are significant at any of the conventional significance levels. The point estimate of the marginal effect of Δlog(IPC_h(-2)) is again negative but the point estimate for the coefficient of Δlog(IPC_h(-3)) is negative as well, which is at odds with our theoretical formulation. Regarding the sign-restricted coefficients, we once again reject that the domestic inflation expectations (output gap) have (has) a lower-than-or-equal-to-zero effect at the 10% significance level and conclude that there exists a positive and significative effect. Compared with the full sample estimation, the point estimate of the coefficient of the expectations is higher which in turn suggests that the expectations channel has gained more relevance after the financial crisis, even when the marginal effect of the output gaps has remained unaltered for both estimation samples. We cannot reject the null hypothesis that there is full indexation at the 10% significance level, which is reflected by a two-sided p-value equal to 0.11 for columns IV and V. However, such type of result is not reflected in column VI and this partly reflects that the adjusted number of observations (135) is considerably lower than the one originally employed. Under such situation, the large-sample distributions might not constitute an acceptable approximation to their finite-sample counterparts. That being said, the results for the post-crisis period should be interpreted with caution.

6 Conclusions

In this paper, we estimated a reduced-form version of the NKPC for the Peruvian economy and the 2002-2019 period. Our empirical evidence supports the argument that the slope of the Phillips curve for Peru has remained stable. At the same time, the expectation channel has gained more relevance in the aftermath of the last financial crisis and this fact is consistent with a lower fraction of producers indexing their prices. Of course, a model-consistent explanation requires an estimation of the model parameters. In this sense, the GMM estimator under structural change by Antoine and Boldea (2018) is particularly promising for both semi-structural and structural estimation.

References

Adolfson, M., S. Laseen, J. Linde, and M. Villani (2007): "Bayesian estimation of an open economy DSGE model with incomplete pass-through," *Journal of International Economics*, 72, 481–511.

Alonso-Carrera, J. and T. Kam (2016): "Anatomizing Incomplete-Markets Small Open Economies: Policy Trade-Offs And Equilibrium Determinacy," *Macroeconomic Dynamics*, 20, 1022–1050.

Andrews, D. W. K. (1999): "Consistent Moment Selection Procedures for Generalized Method of Moments Estimation," *Econometrica*, 67, 543–564.

Anguyo, F. L., R. Gupta, and K. Kotzé (2017): "Monetary Policy, Financial Frictions and Structural Changes: A Markov-Switching DSGE Approach," Working Papers 201748, University of Pretoria, Department of Economics.

Antoine, B. and O. Boldea (2018): "Efficient estimation with time-varying information and the New Keynesian Phillips Curve," *Journal of Econometrics*, 204, 268–300.

Balcilar, M., R. Gupta, and K. Kotzé (2017): "Forecasting South African macroeconomic variables with a Markov-switching small open-economy dynamic stochastic general equilibrium model," *Empirical Economics*, 53, 117–135.

Bouakez, H. and A. Eyquem (2015): "Government spending, monetary policy, and the real exchange rate," *Journal of International Money and Finance*, 56, 178–201.

Cacciatore, M., F. Ghironi, and Y. Lee (2016): "Financial market integration, exchange rate policy, and the dynamics of business and employment in Korea," *Journal of the Japanese and International Economies*, 42, 79–99.

Calvo, G. A. (1983): "Staggered prices in a utility-maximizing framework," *Journal of Monetary Economics*, 12, 383–398.

Campa, J. M. and L. S. Goldberg (2005): "Exchange Rate Pass-Through into Import Prices," *The Review of Economics and Statistics*, 87, 679–690.

Cavaliere, G., D. I. Harvey, S. J. Leybourne, and A. R. Taylor (2011): "Testing For Unit Roots In The Presence Of A Possible Break In Trend And Nonstationary Volatility," *Econometric Theory*, 27, 957–991.

Clarida, R., J. Galí, and M. Gertler (1999): "The Science of Monetary Policy: A New Keynesian Perspective," *Journal of Economic Literature*, 37, 1661–1707.

Craighead, W. D. (2014): "Monetary rules and sectoral unemployment in open economies," *Journal of Macroeconomics*, 40, 277–292.

Davis, J. S., I. Fujiwara, and J. Wang (2017): "Dealing with time-inconsistency: Inflation targeting vs. exchange rate targeting," CAMA Working Papers 2017-54, Centre for Applied Macroeconomic Analysis, Crawford School of Public Policy, The Australian National University.

de Gregorio, J., O. Landerretche, and C. Neilson (2007): "Another Pass-Through Bites the Dust? Oil Prices and Inflation," *Economía Journal*, 0, 155–208.

Debortoli, D. and R. Nunes (2014): "Monetary Regime Switches and Central Bank Preferences," *Journal of Money, Credit and Banking*, 46, 1591–1626.

Dickey, D. A. and W. A. Fuller (1979): "Distribution of the Estimators for Autorregresive Time Series with a Unit Root," *Journal of the American Statistical Association*, 74, 427–31.

Dungey, M., T. Vehbi, and C. Martin (2014): "VAR modelling in the presence of China's rise : an application to the Taiwanese economy," Working Papers 2014-09, University of Tasmania, Tasmanian School of Business and Economics.

Elliott, G., T. J. Rothenberg, and J. H. Stock (1996): "Efficient Tests for an Autoregressive Unit Root," *Econometrica*, 64, 813–836.

Etro, F. and L. Rossi (2015): "Optimal monetary policy under Calvo pricing with Bertrand competition," *Journal of Macroeconomics*, 45, 423–440.

Fasolo, A. M. (2014): "The Ramsey Steady State under Optimal Monetary and Fiscal Policy for Small Open Economies," Working Papers Series 357, Central Bank of Brazil, Research Department.

Galí, J. (2015): *Monetary Policy, Inflation, and the Business Cycle. An Introduction to the New Keynesian Framework and Its Applications*, Princeton University Press, 2nd ed.

Galí, J. and T. Monacelli (2005): "Monetary Policy and Exchange Rate Volatility in a Small Open Economy," *Review of Economic Studies*, 72, 707–734.

Gomez, V. and A. Maravall (1994): "Estimation, Prediction, and Interpolation for Nonstationary Series with the Kalman Filter," *Journal of the American Statistical Association*, 89, 611–624.

Hansen, L. P. (1982): "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica*, 50, 1029–1054.

Justiniano, A. and B. Preston (2010): "Monetary policy and uncertainty in an empirical small open-economy model," *Journal of Applied Econometrics*, 25, 93–128.

Kamber, G., C. McDonald, N. Sander, and K. Theodoridis (2015): "A structural model for policy analysis and forecasting: NZSIM," Reserve Bank of New Zealand Discussion Paper Series DP2015/05, Reserve Bank of New Zealand.

——— (2016): "Modelling the business cycle of a small open economy: The Reserve Bank of New Zealand's DSGE model," *Economic Modelling*, 59, 546–569.

Kavtaradze, L. (2014): "Inflation Dynamics in Georgia," MPRA Paper 59966, University Library of Munich, Germany.

Kichian, M. and F. Rumler (2014): "Forecasting Canadian inflation: A semi-structural NKPC approach," *Economic Modelling*, 43, 183–191.

Kolasa, M. and M. Rubaszek (2018): "Does the foreign sector help forecast domestic variables in DSGE models?" *International Journal of Forecasting*, 34, 809–821.

Kriwoluzky, A., G. J. Müller, and M. Wolf (2015): "Exit Expectations and Debt Crises in Currency Unions," IWH Discussion Papers 18/2015, Halle Institute for Economic Research (IWH).

Lie, D. (2018): "Observed inflation-target adjustments in an estimated DSGE model for Indonesia: Do they matter for aggregate fluctuations?" Working Papers 2018-01, University of Sydney, School of Economics.

MacKinnon, J. G. (1996): "Numerical Distribution Functions for Unit Root and Cointegration Tests," *Journal of Applied Econometrics*, 11, 601–618.

Magnusson, L. M. and S. Mavroeidis (2014): "Identification Using Stability Restrictions," *Econometrica*, 82, 1799–1851.

Monacelli, T. (2005): "Monetary Policy in a Low Pass-through Environment," *Journal of Money, Credit and Banking*, 37, 1047–1066.

Ng, S. and P. Perron (2001): "Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power," *Econometrica*, 69, 1519–1554.

Okano, E. (2014): "How important is fiscal policy cooperation in a currency union?" *Journal of Economic Dynamics and Control*, 38, 266–286.

Perron, P. (1989): "The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis," *Econometrica*, 57, 1361–1401.

Perron, P. and G. Rodríguez (2003): "GLS detrending, efficient unit root tests and structural change," *Journal of Econometrics*, 115, 1–27.

Phillips, A. W. (1958): "The Relation between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom, 1861-1957," *Economica*, 25, 283–299.

Phillips, P. C. B. and P. Perron (1988): "Testing for a Unit Root in Time Series Regression," *Biometrika*, 75, 335–346.

Popescu, I. V. (2014): "The impact of the recent global crisis on the prioritization of central banks final objectives. A structural approach in the context of Central and Eastern European states," *International Journal of Business and Economic Sciences Applied Research (IJBESAR)*, 7, 51–76.

Portillo, R. A. and L.-F. Zanna (2015): "On the First-Round Effects of International Food Price Shocks; the Role of the Asset Market Structure," IMF Working Papers 15/33, International Monetary Fund.

Said, S. E. and D. A. Dickey (1984): "Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order," *Bometrika*, 71, 599–607.

Sánchez, S. C. (2018): "Monetary policy and structural changes in Colombia, 1990-2016: A Markov Switching approach," DOCUMENTOS CEDE 016970, UNIVERSIDAD DE LOS ANDES-CEDE.

Sbordone, A. M. (2005): "Do expected future marginal costs drive inflation dynamics?" *Journal of Monetary Economics*, 52, 1183–1197.

Yeh, K.-c. (2017): "Asset price targeting in an open economy with cognitive limitations: The best for macroeconomic and financial stability?" *The North American Journal of Economics and Finance*, 39, 288–299.

Yun, T. (1996): "Nominal price rigidity, money supply endogeneity, and business cycles," *Journal of Monetary Economics*, 37, 345–370.

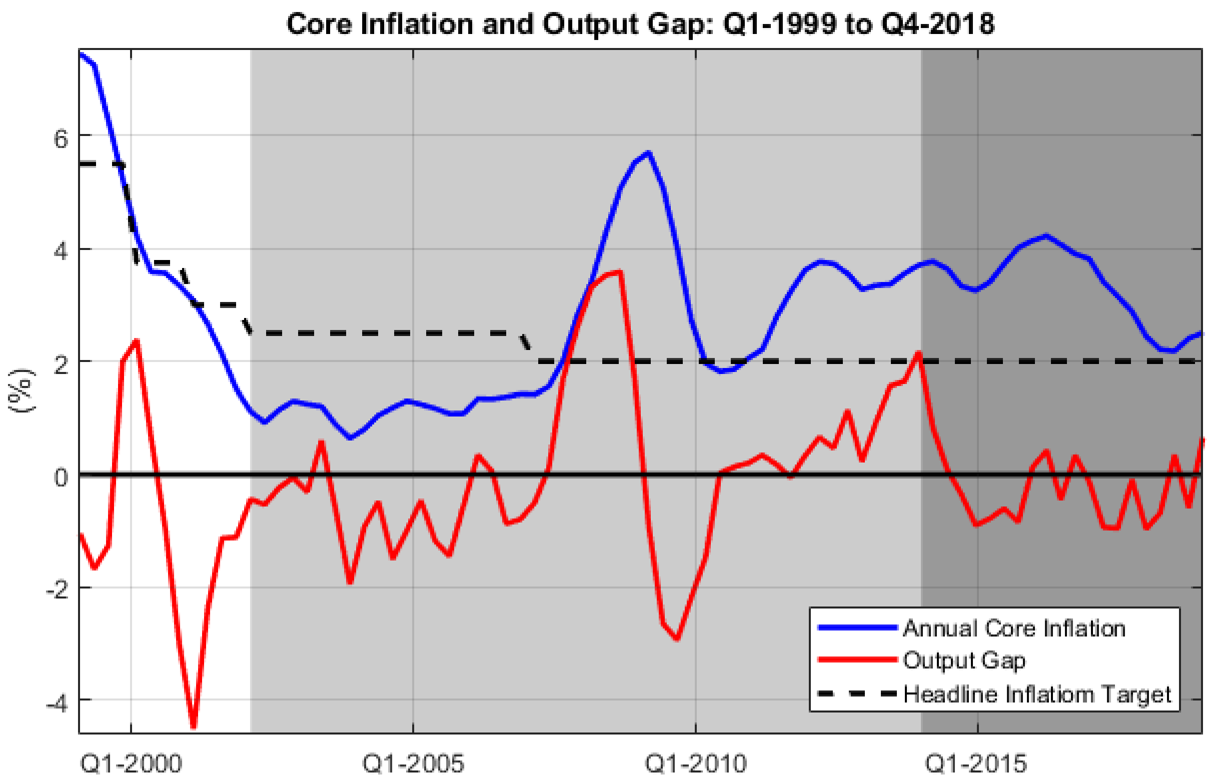

Figure 1: Quarterly Core Inflation and Output Gap

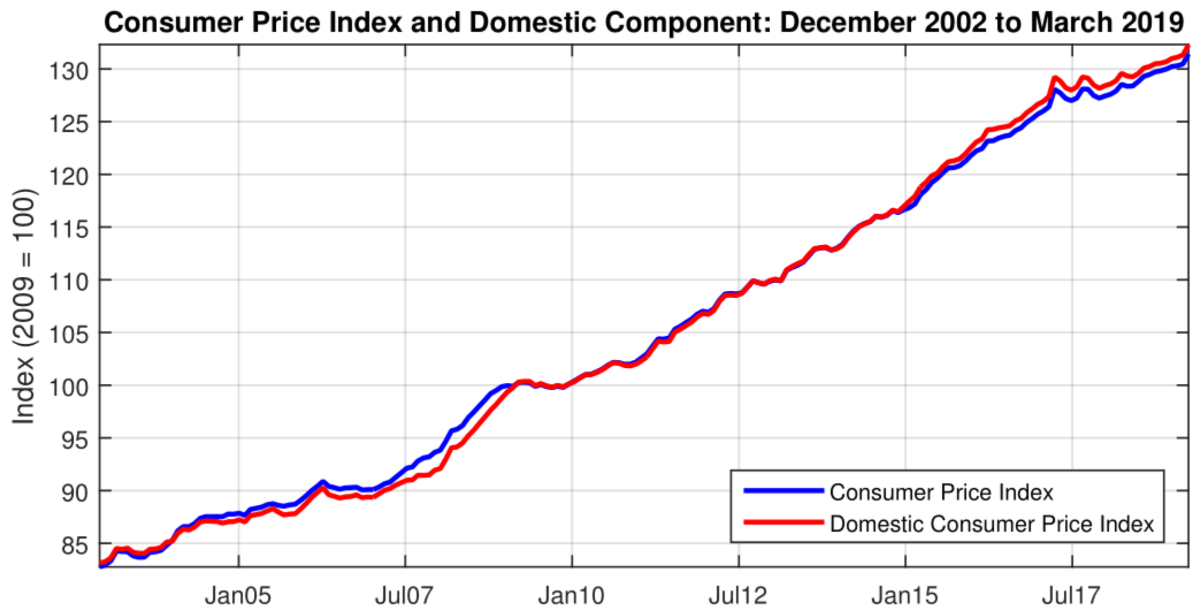

Figure 2: Consumer Price Index and Domestic Component

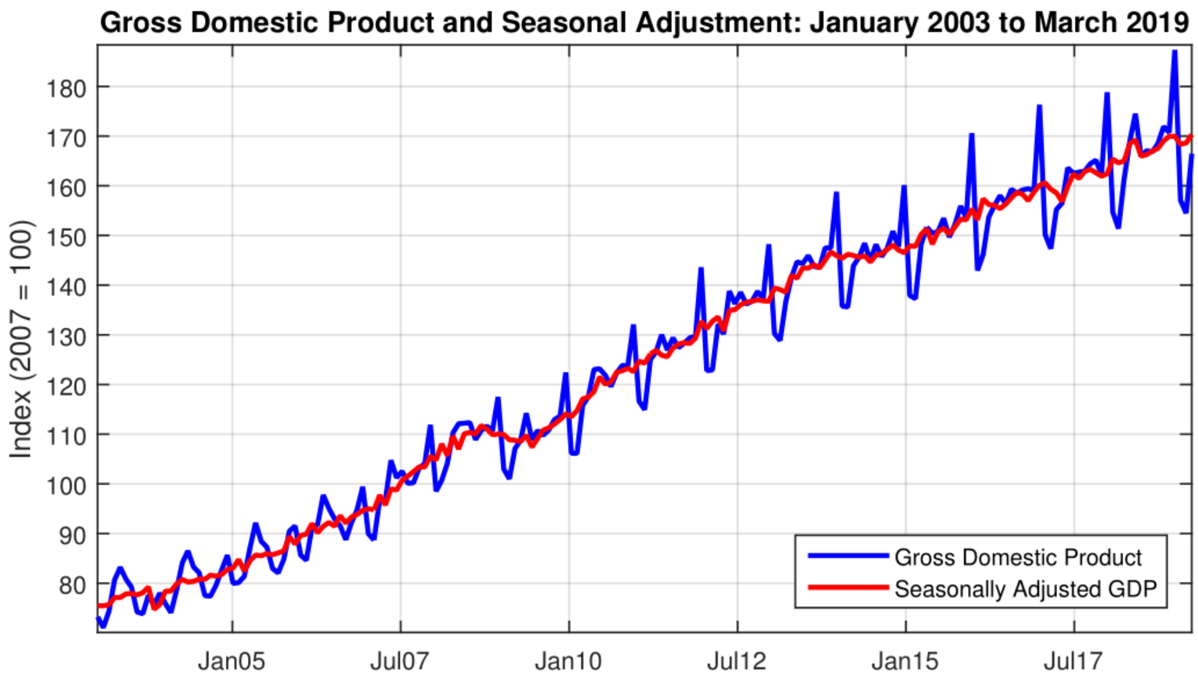

Figure 3: Gross Domestic Product and Seasonal Adjustment

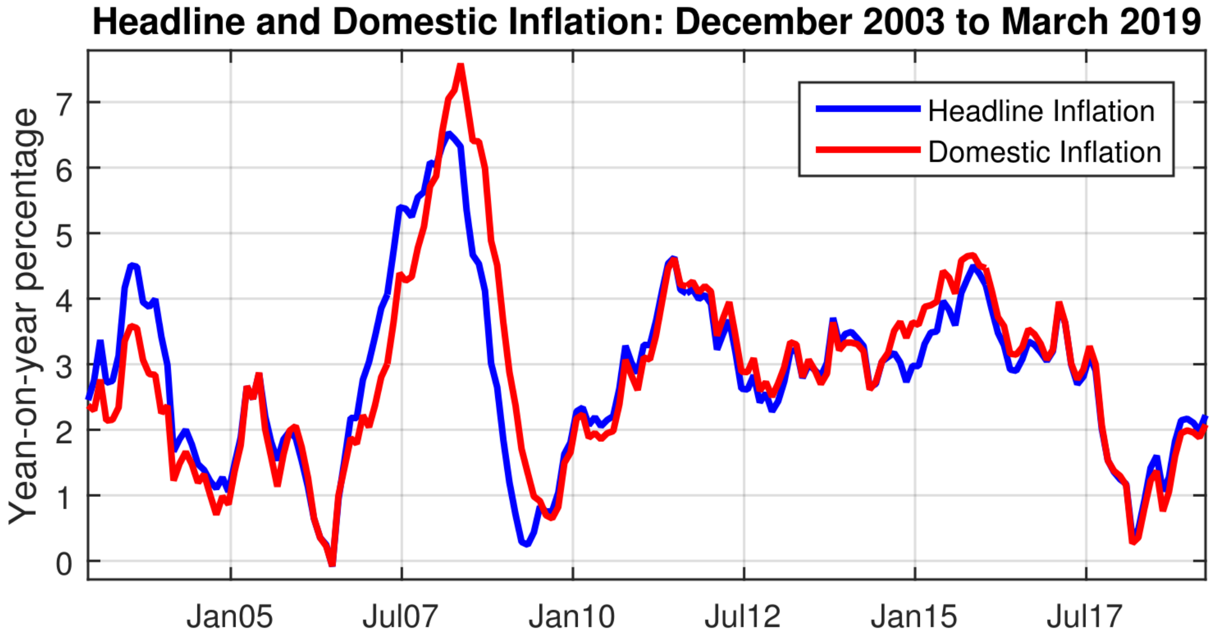

Figure 4: Headline and Domestic Inflation

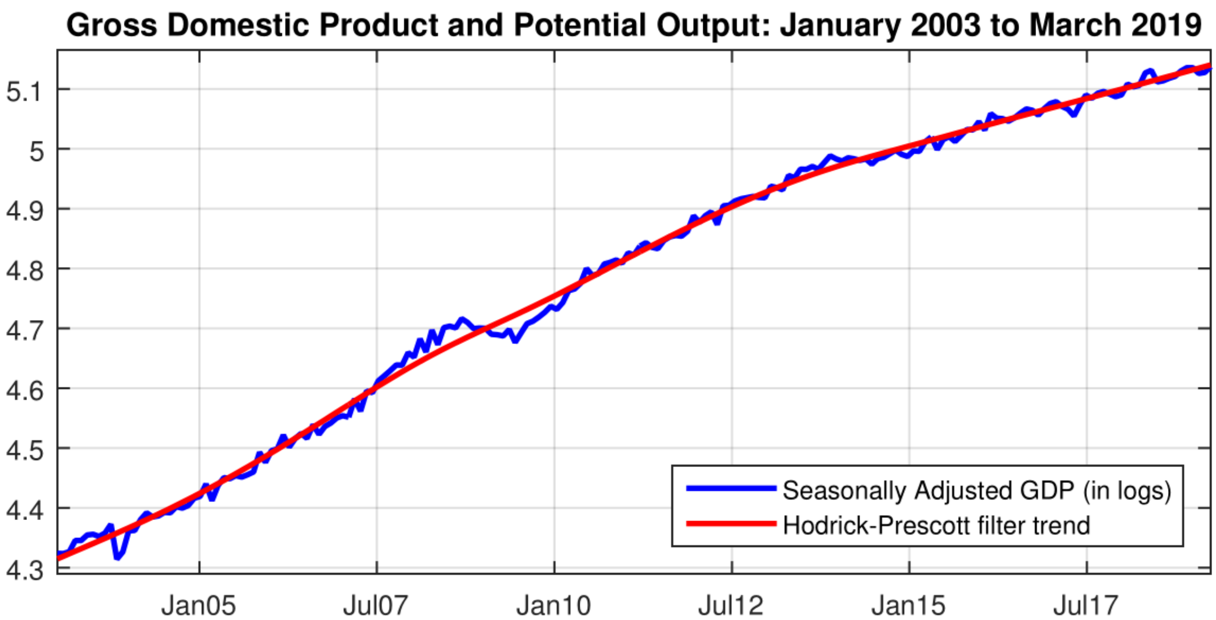

Figure 5: Gross Domestic Product and Potential Output

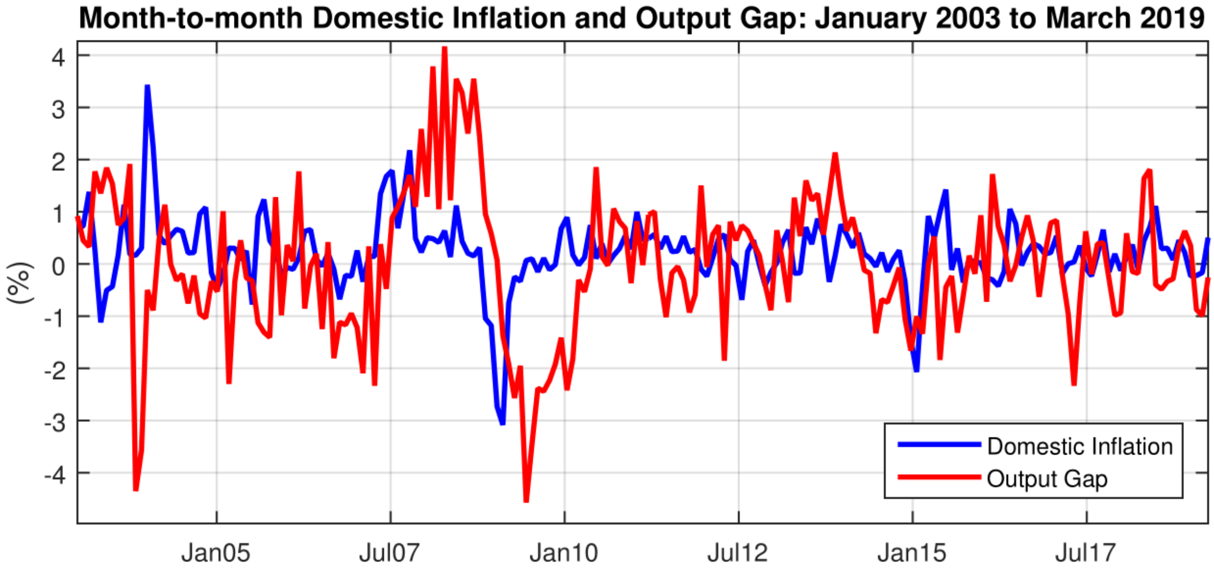

Figure 6: Month to month Domestic Inflation and Output

Table 1: Augmented Dickey-Fuller and Phillips-Perron Unit Root Tests[a]

		No drift nor trend	Drift, no trend	Drift and trend
Augmented Dickey-Fuller Tests				
output_gap		-4.6286***	-4.6167***	-4.6047***
log(IPC_h)		6.3658	0.2604	-2.5142
Critical values[b]	1%	-2.5770	-3.4643	-4.0063
	5%	-1.9425	-2.8764	-3.4333
	10%	-1.6156	-2.5747	-3.1405
Phillips-Perron Tests				
		No drift nor trend	Drift, no trend	Drift and trend
output_gap		-7.9415***	-7.9244***	-7.9066***
log(IPC_h)		8.5337	0.1812	-2.2680
Critical values[b]	1%	-2.5769	-3.4641	-4.0061
	5%	-1.9425	-2.8763	-3.4332
	10%	-1.6156	-2.5747	-3.1404

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] MacKinnon (1996) one-sided p-values.

Table 2: Elliott-Rothenberg-Stock Unit Root Tests[a]

		Intercept
output_gap		1.0243***
log(IPC_h)		1218.5560
Asymptotic critical values[b]	1%	1.9120
	5%	3.1670
	10%	4.3320
		Trend and Intercept
output_gap		2.8870***
log(IPC_h)		15.0534
Asymptotic critical values[b]	1%	4.0605
	5%	5.6590
	10%	6.8565

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] Elliott et al. (1996, Table 1).

Table 3: Elliott-Rothenberg-Stock DF-GLS Unit Root Tests[a]

		Intercept
output_gap		-3.8804***
log(IPC_h)		4.2886
Asymptotic critical values[b]	1%	-2.5770
	5%	-1.9425
	10%	-1.6156
		Trend and Intercept
output_gap		-4.3942***
log(IPC_h)		-1.8133
Asymptotic critical values[b]	1%	-3.4684
	5%	-2.9370
	10%	-2.6470

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.
[b] Elliott et al. (1996, Table 1).

Table 4: Ng-Perron Unit Root Tests[a]

		Intercept			
		MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MPT^{GLS}
output_gap		-25.9777***	-3.5904***	0.1382***	0.9886***
log(IPC_h)		1.6456	4.8807	2.9660	639.6520
Asymptotic critical values[b]	1%	-13.8000	-2.5800	0.1740	1.7800
	5%	-8.1000	-1.9800	0.2330	3.1700
	10%	-5.7000	-1.6200	0.2750	4.4500
		Trend and Intercept			
		MZ_{α}^{GLS}	MZ_{t}^{GLS}	MSB^{GLS}	MPT^{GLS}
output_gap		-31.4988***	-3.9684***	0.1260***	2.8938***
log(IPC_h)		-6.4635	-1.7976	0.2781	14.0984
Asymptotic critical values[b]	1%	-23.8000	-3.4200	0.1430	4.0300
	5%	-17.3000	-2.9100	0.1680	5.4800
	10%	-14.2000	-2.6200	0.1850	6.6700

[a] *, ** and *** indicate rejection of the $I(1)$ null hypothesis at the 10%, 5% and 1% level of significance, respectively. Modified or M-tests are described in Ng and Perron (2001). For the the case of the MZ_{α}^{GLS}, MZ_{t}^{GLS} and MSB^{GLS} tests, a statistic lower than the critical value leads to a rejection of the $I(1)$ null hypothesis.

[b] Ng and Perron (2001, Table 1).

Table 5: Perron-Rodríguez Unit Root Tests[a]

		sup MZ_{α}^{GLS}	sup MZ_{t}^{GLS}	sup MSB^{GLS}
output_gap		-27.0414***	-3.6770***	0.1360**
log(IPC_h)		-15.4456	-2.6837	0.1738
Critical values[b]	1%	-27.0000	-3.6600	0.1340
	5%	-22.9000	-3.3500	0.1450
	10%	-20.7000	-3.1900	0.1540

[a] *, ** and *** indicate rejection of the $I(1)$ null hypothesis at the 10%, 5% and 1% level of significance, respectively. Modified or M-tests under structural change are described in Perron and Rodríguez (2003). In the case of sup MZ_{α}^{GLS}, sup MZ_{t}^{GLS} and sup MSB^{GLS} tests, a statistic lower than the critical value leads to a rejection of the $I(1)$ null hypothesis.

[b] Perron and Rodríguez (2003), Table 2.

Table 6: Cavaliere-Harvey-Leybourne-Taylor Unit Root Tests[a]

		MZ_{α}	MZ_t	MSB	$t(\bar{\tau})$
output_gap		-25.4480**	-3.5670**	0.1400**	-3.7580**
Critical values[b]	(5%)	-16.2190	-2.8150	0.1730	-2.9110
log(IPC_h)		-15.7110	-2.7100	0.1720	-2.7330
Critical values[b]	(5%)	-23.1760	-3.3930	0.1460	-3.6380

[a] *, ** and *** indicate rejection of the unit root hypothesis at the 10%, 5% and 1% level of significance, respectively.

[b] Critical values are computed via the bootstrap algorithm by Cavaliere et al. (2011, Section 4).

Table 7: Estimation results[a]

Equation	(I)	(II)	(III)	(IV)	(V)	(VI)
Estimation method	GMM	GMM	GMM	GMM	GMM	GMM
Period	Jan2002-Mar2019	Jan2002-Mar2019	Jan2002-Mar2019	Jan2008-Mar2019	Jan2008-Mar2019	Jan2008-Mar2019
Dependent variable	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)	Δlog(IPC_h)
Constant	0.1613	0.1509	0.1654	0.1921	0.1921	0.1994*
	(0.1016)	(0.1046)	(0.1039)	(0.1306)	(0.1321)	(0.1024)
Δlog(IPC_h(-1))	0.3237***	0.3205***	0.3261***	0.1692***	0.1680***	0.1977***
	(0.1048)	(0.1050)	(0.1062)	(0.0954)	(0.0963)	(0.0864)
Δlog(IPC_h(-2))	-0.5265	-0.5183	-0.5587	-0.0508	-0.0142	-0.0700
	(0.3304)	(0.3312)	(0.3525)	(0.1112)	(0.1113)	(0.0869)
Δlog(IPC_h(-3))	0.1027	0.1008	0.1119	-0.5864	-0.6320	-0.4099
	(0.0916)	(0.0919)	(0.1000)	(0.3356)	(0.3279)	(0.2799)
Δlog(IPC_h(+1))	0.4386+	0.4710+	0.4369+	0.6887+	0.6931+	0.5084+
	(0.3213)	(0.3310)	(0.3274)	(0.4270)	(0.4285)	(0.3463)
One-sided p-value[b]	0.0869	0.0783	0.0919	0.0546	0.0541	0.0722
output_gap	0.0738+	0.0720+	0.0757+	0.0834+	0.0846+	0.0628+
	(0.0473)	(0.0476)	(0.0483)	(0.0586)	(0.0592)	(0.0444)
One-sided p-value[c]	0.0604	0.0661	0.0595	0.0786	0.0777	0.0799
P-value (full indexation)[d]	0.1326	0.1623	0.1311	0.1088	0.1094	0.0379
No. of observations	189	189	183	135	135	135
No. of instruments	12	11	11	13	12	12
Sargan's J-statistic	0.5903	0.4059	0.5204	4.7417	3.0685	3.9533
Prob(J-statistic)	0.9966	0.9952	0.9914	0.6914	0.8002	0.6830
GMM-BIC	-30.8602	-25.8029	-25.6883	-29.5952	-26.3631	-25.4784
GMM-AIC	-11.4097	-9.5941	-9.4796	-9.2583	-8.9315	-8.0467
GMM-HQIC	-19.3889	-16.2435	-16.1290	-17.6339	-16.1106	-15.2259

[a] *, ** and *** indicate rejection of the null hypothesis of a zero coefficient at the 10%, 5% and 1% level of significance, respectively. Also, +, ++ and +++ indicate rejection of the null hypothesis of a lower-than-or-equal-to-zero coefficient at the 10%, 5% and 1% level of significance, respectively. Reported moment-selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are computed as described by Andrews (1999, Section 3).

[b] One-sided p-value for testing $H_0 : c_5 \leq 0$ against $H_1 : c_5 > 0$ in (8).

[c] One-sided p-value for testing $H_0 : c_6 \leq 0$ against $H_1 : c_6 > 0$ in (8).

[d] Two-sided p-value for testing $H_0 : c_2 + c_3 + c_4 + c_5 = 1$ against $H_1 : c_2 + c_3 + c_4 + c_5 \neq 1$ in (8).

Table 8: Instruments[a]

Equation	(I)	(II)	(III)	(IV)	(V)	(VI)
Constant	Yes	Yes	Yes	Yes	Yes	Yes
Δlog(IPC_h(-1))	Yes	Yes	Yes	Yes	Yes	Yes
Δlog(IPC_h(-2))	No	No	No	Yes	Yes	Yes
Δlog(IPC_h(-3))	Yes	Yes	Yes	No	No	No
Δlog(IPC_h(-4))	No	No	No	Yes	Yes	Yes
Δlog(IPC_h(-5))	Yes	Yes	No	Yes	Yes	Yes
Δlog(IPC_h(-6))	Yes	No	Yes	Yes	Yes	Yes
Δlog(IPC_h(-7))	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-1)	Yes	Yes	Yes	No	No	Yes
output_gap(-2)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-3)	Yes	Yes	Yes	Yes	Yes	No
output_gap(-4)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-5)	Yes	Yes	Yes	Yes	Yes	Yes
output_gap(-6)	Yes	Yes	Yes	Yes	No	No
output_gap(-7)	No	No	No	Yes	Yes	Yes
GMM BIC	-30.8602	-25.8029	-25.6883	-29.5952	-26.3631	-25.4784
GMM AIC	-11.4097	-9.5941	-9.4796	-9.2583	-8.9315	-8.0467
GMM HQIC	-19.3889	-16.2435	-16.1290	-17.6339	-16.1106	-15.2259

[a] Reported moment-selection criteria (GMM-BIC, GMM-AIC and GMM-HQIC) are computed as described by Andrews (1999, Section 3).